EMU and intra-European trade.
Long-run evidence using gravity equations.

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Abstract. In this article we present evidence of the long-run effect of the euro on trade for the twelve initial EMU countries for the period 1967-2008 from a double perspective. First, we pool all the bilateral combinations of trade flows among the EMU countries in a panel cointegration gravity specification. Second, we estimate a gravity equation for each of the EMU-members vis-à-vis the other eleven partners. Whereas the joint gravity equation provides evidence on the aggregate effect of the euro on intra-European trade, by isolating the individual countries we assess which of the member countries have obtained a larger benefit from the euro. Moreover, this strategy permits to check the robustness of the aggregate results and to find possible asymmetries. From an econometric point of view, we apply panel cointegration techniques based on factor models that account for cross-dependence and structural breaks.

***Keywords*:** Gravity models; trade; euro; panel cointegration; structural breaks, cross-section dependence.

***JEL classification*:** C12, C22, F15, F10.

1. Introduction.

The effect that a currency union has on trade has been largely explored in the literature. Rose (2000) is one of the most cited articles in this field, and his prediction of a tripling of trade for a country when it joins a currency union has been revisited several times. Moreover, the creation of EMU have provided to researchers a natural experiment to further investigate on this effect, thus renewing the debate and leading to improvements in both the specification and estimation of the gravity equation. Though initial estimates where found to be quite high, ranging from the approximate 2% in Glick and Rose (2002) to the 27% in Barr et al. (2003), more recent literature has considerably reduced this effect. On the other hand, as Camarero et al. (2011) claim, the creation of the EMU is better interpreted as a continuation, or culmination, of a series of policy changes that have led over the last four decades to greater economic integration among the countries that now constitute the EMU, the euro having just a residual effect. Other articles supporting this hypothesis are with Bun and Klaasen (2007), Fidmurc (2009), Gengenbach (2009) and Berger and Nitsch (2008). Hiller and Kruse (2010) provide an analysis of this integration process, finding the most relevant dates in the integration process for each one of the EMU countries.

A logical consequence of this fact is that, given that the Euro is a long-run process, long run estimation methods are more appropriated to measure the Euro effect and, hence, the nonstationarity of variables or the existence of cointegration relationships among the variables should be controlled for when estimating the gravity equation to avoid biases and inconsistencies. For that reason, the use of cointegration techniques and the inclusion of time trends in the specification have appeared in recent trade literature, contributing to better estimate the euro effect. Gengenbach (2009), Faruquee (2004) or Camarero et al. (2011) introduces the use of cointegration techniques and find estimates that range from 0.6% to 8%.

However, there is still an important caveat in the literature. Frequently the cointegrating relationship is assumed to be stable. Nevertheless, failure to account for the existence of changes in the cointegration relationship and/or the deterministic components affects inference on cointegration analysis, thus leading to wrong conclusions. The standard tests may not reject the null hypothesis of no cointegration when it is false, thus reducing the power of the test. As far as we know, Camarero et al. (2011) and Mancini-Griffoli and Pauwels (2006) are the only articles allowing for the possibility of structural breaks in data when estimating the gravity equation using cointegration techniques. In the case of Mancini-Griffoli and Pauwels (2006) the date of the break is found in the first quarter of 1999 and three alternative specifications of the gravity equation are estimated using DOLS and an ECM. However, these estimators do not correct for cross-section dependence. Since the Pesaran CD statistic reveals the existence of these dependencies, we claim that robust estimators should be employed. In Camarero et al. (2011), Banerjee and Carrión-i-Silvestre (2010) test is employed to properly specify the equation and the break is found to happen in 1987.

Finally, there is little evidence on the asymmetric effect of the Euro on the different Euro members using cointegration techniques. Faruquee (2004) makes a comparision of the effect of EMU on its member by interacting country dummies with EMU variable. His results show that Norway and Spain are the countries that have obtained greater benefits of joining EMU, while Ireland, Finland and Portugal are the countries with lower benefits. Dwane et al. (2011) perform an analysis of this effect, but focusing only on Irish trade. However, in both cases the possibility of breaks is ignored and cross section dependencies are not modeled. In this article, we investigate the aggregate effect of the Euro on intra-European trade as well as the specific effect on each one of its members in a panel cointegration framework, allowing for structural breaks in the specification. We employ Bai et al. (2009) CUP estimator, which is consistent in the presence of cross section dependencies, and use a more homogeneous sample, which is more appropriate when the date of the break is unique. To the best of our knowledge, estimators robust to cross section dependencies and structural breaks have never been applied before to the estimation of the Euro effect.

The remainder of the article is organized as follows. In section 2 we describe the data and the variables utilized in the analysis, as well as the methodology and tests employed. In section 3 we present the results for the EMU as a whole. In section 4, a gravity equation for each of the EMU-members vis-à-vis the other eleven partners is estimated and the Euro effect is studied country by country. Finally, section 5 concludes.

2. Data, methodology and empirical results.

2.1. Data and model

We include in our study all the countries that joined EMU on 1999 and Greece, which joined EMU in 2001. Belgium and Luxembourg are included as a unique area, so the total number of countries included is 11[1]. The sample contains annual data and covers the period 1967-2008. Hence, we have a balanced panel with dimension N=110 (all possible bilateral combinations of countries) and T= 42. The total number of observations is NT =4,620. Although the number of years available was higher, we have opted by restrict our sample to this period, in order to exclude the effects of the financial crisis that started in 2008.

[1] Austria, Belgium and Luxembourg, Finland, France, Germany, Greece, Ireland, Italy, Netherlands, Portugal, Spain.

According to Baldwin and Taglioni (2006) critiques, the variables are introduced in real terms, and trade is measured as the sum of the logarithm of imports and exports.

We use the specification of the gravity equation in Camarero et al. (2011) which is similar to the traditional specification from the recent literature on the euro effect using nonstationary panels:

$$TRADE_{ijt} = \beta_1 GDP_{ijt} + \beta_2 GDPCAP_{ijt} + \delta_1 EURO_{ijt} + \delta_2 RTA_{ijt} + \eta_{ij} + \tau_{ij}\, t + \lambda_t + \varepsilon_{ijt}$$

The dependent variable in our gravity model is $TRADE_{ijt}$, as the sum of the logarithm of imports and exports in nominal terms. GDP_{ijt} and $GDPCAP_{ijt}$ are the bilateral products of nominal GDP and GDP per capita in each pair of countries, coming from CHELEM database. RTA_{ijt} is a dummy variable that takes value one when both countries belong to a Regional Trade Agreement. $EURO_{ijt}$ is our variable of interest; it takes value one when both countries use the Euro.

The purpose is to isolate the effects of EMU on trade trying to control for other factors that may have an influence on trade flows but are not related to the monetary union

2.2. Panel unit root tests and cross-section dependence

In this section we test the presence of cross-section dependence among the series using the test statistic by Pesaran (2004) to assess whether the time series in the panel are cross-section independent. Under the null hypothesis of cross section independence the CD statistic of Pesaran (2004) converges to the standard normal distribution. The results in Table 1 show that the Pesaran's CD statistic strongly rejects the null hypothesis of independence, so that cross-section dependence has to be considered when computing the panel data statistics if misleading conclusions are to be avoided. Note that, according to Pesaran (2004) the CD test is valid for N and T tending to ∞ in any order and that it is particularly useful for panels with small T and large N. Moreover, this test is also robust to possible structural breaks, which makes it especially suitable for our study.

Once we have found the presence of dependence in the variables, we study the order of integration of the variables. Bai and Ng (2004) is a suitable approach when cross-correlation is pervasive, as in this case. Furthermore, this approach controls for cross-section dependence given by cross-cointegration relationships, potentially possible among our group of countries and variables — see Banerjee et al. (2004). The Bai and Ng (2004) approach decomposes the $Y_{i,t}$, time series as follows:

$$Y_{i,t} = D_{i,t} + F_t^{'} \pi_i + e_{i,t}$$

with t = 1, . . . , T , i = 1, . . . , N, where $D_{i,t}$ denotes the deterministic part of the model — either a constant or a linear time trend — F_t is a (r x1)-vector that accounts for the common factors that are present in the panel, and $e_{i,t}$ is the idiosyncratic disturbance term, which is assumed to be cross-section independent. Unobserved common factors and idiosyncratic disturbance terms are estimated using principal components on the first difference model. For the estimated idiosyncratic component, they propose an ADF test for individual unit roots and a Fisher-type test for the pooled unit root hypothesis ($P_{\hat{e}}$), which has a standard normal distribution. The estimation of the number of common factors is obtained using the panel BIC information criterion as suggested by Bai and Ng (2002), with a maximum of six common factors. Bai and Ng (2004) propose several tests to select the number of independent stochastic trends, k_1 in the estimated common factors, $\hat{F}_t$. If a single common factor is estimated, they recommend an ADF test whereas if several common factors are obtained, they propose an iterative procedure to select k_1: two modified Q statistics (MQ_c and MQ_f), that use a non-parametric and a parametric correction respectively to account for additional serial correlation. Both statistics have a non-standard limiting distribution. They test the hypothesis of $k_1 = m$ against the alternative $k_1 < m$ for m starting from $\hat{k}$. The procedure ends if at any step $k_1 = m$ cannot be rejected. The results from the application of the Bai and Ng (2004) statistics are summarized in Table 2.

2.3. Panel cointegration test

Banerjee and Carrion-i-Silvestre (2010) propose panel tests for the null hypothesis of no cointegration allowing for breaks both in the deterministic components and in the cointegrating vector. In addition, they tackle cross-section dependence using factor models.

Let $Y_{i,t} = (y_{i,t}, x_{i,t}^{'})$ be a $(m \times 1)$-vector of non-stationary stochastic process whose elements are individually $I(1)$ with the following Data Generating Process (DGP):

$$y_{i,t} = D_{i,t} + x_{i,t}^{'}\delta_{i,t} + u_{i,t} \tag{1}$$

The general functional form for the deterministic term $D_{i,t}$ is given by:

$$D_{i,t} = \mu_i + \beta_i t + \sum_{j=1}^{m_i}\theta_{i,j} DU_{i,j,t} + \sum_{j=1}^{m_i}\gamma_{i,j} DT_{i,j,t}, \tag{2}$$

where $DU_{i,j,t}$ =1 and $DT_{i,j,t}$ = (t - $T_{i,t}^{b}$) for $t > T_{i,t}^{b}$ and 0 otherwise, $T_{i,t}^{b} = \lambda_{i,j}^{b} T$ denotes the timing of the j-th break, $j = 1,\ldots, m_i$, for the i-th unit, $I = 1,\ldots, N$, $\lambda_{i,j}^{b} T \in \Lambda$, being Λ a closed subset of (0,1).

Banerjee and Carrion-i-Silvestre (2010) propose six different model specifications:

Model 1. Constant term, no linear trend - $\boldsymbol{\theta_{ij} = \beta_i = \gamma_{i,j} = 0}$ $\forall i, j$ in (2) – and constant cointegrating vector.

Model 2. Stable trend - $\boldsymbol{\theta_{ij} = 0; \beta_i \neq 0}$ $\forall i$ and $\boldsymbol{\gamma_{i,j} = 0}$ $\forall i, j$ in (2) - and constant cointegrating vector.

Model 3. Constant term with shifts; stable trend - $\boldsymbol{\Theta_{ij} \neq 0; \beta_i \neq 0; \gamma_{i,j} = 0}$ $\forall i, j$ (2) – and constant cointegrating vector. The model considers multiple level shifts.

Model 4. Constant term, trend and changes in trend, **-** $\boldsymbol{\Theta_{ij} = 0; \beta_i \neq \gamma_{i,j} \neq 0}$ $\forall i, j$ in (2) – and constant cointegrating vector. The model considers multiple trend shifts.

Model 5. Changes in constant and trend $\boldsymbol{\Theta_{ij} \neq 0; \beta_i \neq 0}$ $\forall i$ and $\boldsymbol{\gamma_{i,j} \neq 0}$ $\forall i, j$ in (2) – and constant cointegrating vector. The model considers multiple trend and level shifts.

Model 6. No trend, constant term with shifts $\boldsymbol{\Theta_{ij} \neq 0; \beta_i = 0}$ $\forall i, j$ -and changes in the cointegrating vector.

Model 7. Constant term, trend; changes in the level $\boldsymbol{\Theta_{ij} \neq 0; \beta_i \neq 0}$-and changes in the cointegrating vector.

Model 8. Constant term, trend; changes in the level and the trend $\boldsymbol{\Theta_{ij} \neq 0; \beta_i \neq 0}$ and $\boldsymbol{\gamma_{i,j} \neq 0}$ $\forall i, j$ -and changes in the cointegrating vector

The common factors are estimated following the method proposed by Bai and Ng (2004). They first compute the first difference of the model; then, they take the orthogonal projections and estimate the common factors and the factor loadings using principal components.

In any of these specifications, Banerjee and Carrion-i-Silvestre (2010) recover the idiosyncratic disturbance terms ($\tilde{e}_{i,t}$) through cumulation of the estimated residuals and propose testing for the null of no cointegration against the alternative of cointegration with break using the ADF statistic.

The null hypothesis of a unit root can be tested using the pseudo t-ratio $t^{j}_{\tilde{e}_i}(\lambda_i)$, $j = c, \tau, \gamma$. The models that do not include a time trend (Models 1 and 6) are denoted by c. Those that include a linear time trend with stable trend (Models 2, 3 and 7) are denoted by τ and, finally, γ refers to the models with a time trend with changing trend (Models 4, 5 and 8).

When common (homogeneous) structural breaks are imposed to all the units of the panel (although with different magnitudes), we can compute the statistic for the break dates, where the break dates are the same for each unit, using the idiosyncratic disturbance terms[2].

In Table 3 we present the results of the tests for non-cointegration Z_j^* for the model with homogeneous structural breaks for the six potential specifications discussed above. Using the BIC information criterion, we choose model 5, that contains a constant and a trend and the structural break affects them both simultaneously. Using again the BIC information criterion, we find three factors in the panel. In order to test for non-cointegration, we apply the statistics based on the accumulated idiosyncratic components, Z_j^*. We present the tests for all possible model specifications. With all of them the null hypothesis of non-cointegration can be rejected. The break is found to happen in 1986, the year before in which the Single European Act was signed.

Given that the existence of cointegration relationships is unambiguous, the next step of the anaysis is to estimate the long-run relationship in the form of a gravity equation. For this purpose, in the next section we will employ consistent techniques proposed by Bai et al (2009) and Pesaran (2006).

3. Estimation of the gravity equation for the EMU

Bai et al. (2009) consider the problem of estimating the cointegrating vector in a cointegrated panel data model with non-stationary common factors. The presence of

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[2] As described in equations (2) and (3), a heterogeneous version of the test is also possible, authough the homogeneous case is the more adequate for the particular case of the gravity model and the estimation of the parameters in the long-run relationship.

common sources of non-stationarity leads naturally to the concept of cointegration. In addition, by putting a factor structure one can deal with other sources of correlation and with large panels, as it is our case.

Bai et al. (2009) treat the common $I(1)$ variables as parameters. These are estimated jointly with the common slope coefficients β using an iterated procedure. Although this procedure yields a consistent estimator of β, the estimator is asymptotically biased. To account for this bias, the authors construct two estimators that deal with endogeneity and serial correlation and re-center the limiting distribution around zero. The first one, CupBC, estimates the asymptotic bias directly. The second, denoted CupFM, modifies the data so that the limiting distribution does not depend on nuisance parameters. Both are "continuously-updated" (CUP) procedures and require iteration till convergence. The estimators are $\sqrt{nT}$ consistent and enable the use of standard tests for inference. Finally, the approach is robust to mixed $I(1)/I(0)$ factors as well as mixed $I(1)/I(0)$ regressors.

Bai et al. (2009) consider the following model:

$$y_{it} = x_{it}'\beta + e_{it}$$

where for $i = 1,\dots,n$, $t=1,\dots,T$, y_{it} is a scalar,

$$x_{it} = x_{it-1} + \varepsilon_{it}$$

x_{it} is a set of k non-stationary regressors, β is a $k \times 1$ vector of the common slope parameters, and e_{it} is the regression error. They assume that e_{it} is stationary and *iid* across i. The pooled least squares estimator of β is as follows:

$$\hat{\beta}_{LS} = \left(\sum_{i=1}^{n} \sum_{t=1}^{T} x_{it} x_{it}' \right)^{-1} \sum_{i=1}^{n} \sum_{t=1}^{T} x_{it} y_{it}$$

Although his estimator is, in general, T consistent, there is an asymptotic bias due to the long-run correlation between e_{it} and ε_{it}. This bias can be estimated and a panel fully-modified estimator can be developed as in Phillips and Hansen (1990) to achieve $\sqrt{nT}$

consistency and asymptotic normality. In addition, they model cross-section dependence by imposing a factor structure on e_{it}:

$$e_{it} = \lambda_i' F_t + u_{it}$$

where F_{it} is an $r \times 1$ vector of latent common factors, λ_i is an $r \times 1$ vector of factor loadings and u_{it} is the idiosyncratic error. If both F_t and u_{it} are stationary, then e_{it} is also stationary. In this case, a consistent estimator of the regression coefficients can still be obtained even when the cross-section dependence is ignored. Bai and Ng (2006) considered a two-step fully-modified estimator (2sFM).

It is crutial to note that when F_t is $I(1)$, if $F_t = F_{t-1} + \eta_t$, then e_{it} is I(1) and the pooled OLS is not consistent. This is why Bai et al. (2009) develop the case of non-stationary common factors, aiming at achieving consistent estimators.

When the common factor F_t is observed, they propose what can be considered the panel version of the Phillips and Hansen (1990) statistic, a linear estimator that they call $\tilde{\beta}_{LSFM}$ and the bias corrected version that is identical. The estimators are consistent and the limiting distributions are normal.

However, in the majority of the cases, the factors F_t are unobserved. In this case, the LSFM estimator is infeasible. Thus, F_t should be estimated along with β by minimizing the objective function, the unobserved quantities can be replaced by initial estimates and iterate until convergence through the continuously-updated estimator (CUP) for (β, F), defined as $(\hat{\beta}_{Cup}, \hat{F}_{Cup}) = \underset{\beta, F}{\arg\min}\, S_{nT}(\beta, F)$. The estimator $\hat{\beta}_{Cup}$ is consistent for β, although it still has a bias derived from having to estimate F_t. The authors correct this bias using two fully-modified estimators. The first one directly corrects the bias of $\hat{\beta}_{Cup}$ and is denoted $\hat{\beta}_{CupBC}$.

The second one makes the correction in each iteration and is denoted $\hat{\beta}_{CupFM}$.

We present in Table 4 the results of the CUP estimation using the methodology of Bai et al. (2009). We have based our estimation on the results previously obtained using the Banerjee and Carrión-i-Silvestre (2010) tests concerning not only the cointegration tests, but also the deterministic specification of the chosen model. Bai et al. (2009) consider extensions of their estimators when the assumptions about the deterministic components are relaxed. In order to account for the existence of incidental trends (intercept and/or trend), they define accordingly the projection matrix M considered above for demeaned and/or detrended variables. We concentrate the deterministic components before we estimate the long-run parameters. Among those deterministic components we have also included the common structural breaks[3].

However, though the CUP estimators are consistent, recent literature has focused on Pesaran (2006) CCE estimator, in which the common factors are approximated using the cross section averages of each of the variables to capture the common factors without requiring the estimation of their number.

4. Evidence on individual countries

Our sample includes twelve EMU countries which are quite similar and hence the homogeneity of the sample is enough to found a reasonable break common to all of them. However, in this section we give a further step in the determination of the euro effect and repeat the previous analysis for each one of the individual countries with the purpose of eliminating all the possible heterogeneity thus allowing to check the robustness of the

[3] Note that this implies that in the model specification of the gravity equation in expression (1) above, we have filtered the three variables (trade, GDP and GDP per capita) of the deterministic components and the structural breaks, with the exception of the dummies RTA and EMU.

aggregate results and to find possible asymmetries. This allows us to introduce a specific break for each country. Table 5 show the result of this analysis. In most cases the euro effect is found to be positive though small. When we apply consistent estimation techniques, the results are negative for Greece and Italy when using CUP and for Finland, France and Ireland when using CCE.

5. Summary and concluding remarks

In this paper we contribute to the existent literature concerning the euro effect with the application of an estimation method that is consistent in the presence of nonstationarities and dependencies in the data. We use a data set that includes 12 EMU countries from 1967 to 2008 and estimate a gravity equation through a cointegration approach fully allowing for cross-section dependence. The analysis consists of three steps. First, unit root tests for cross-sectionally dependent panels are applied. Second, the existence of a cointegration relationship among the variables of a proper specification of the gravity equation is tested. In this exercise we account both for dependence in the cross-section dimension and discontinuities in the deterministic and the cointegrating vector in the time dimension. Third, three consistent estimation methods are used to estimate the long-run relationships; the first two, CUP-BC and CUP-FM, model the dependencies in data using common factors, whereas the third, the Pesaran (2006) CCE estimator, proxies the dependencies using cross-section averages.

Our specification allows for cross-sections dependence and structural breaks in the time domain as well as nonstationarities in the variables. Our results reinforce those in Camarero et al. (2011), in which it is argued that the creation of the EMU is best interpreted as a continuation, or culmination, of a series of policy changes that have led over the last four decades to greater economic integration among the countries that now constitute the EMU. We find strong evidence of a gradual increase in trade intensity between European countries as well as pervasive cross section dependence. Once we control for both, dependence and this trend in trade integration, the effect of the formation of the EMU fades out in line with most recent empirical literature. In the aggregated case, the break is found to happen in 1986. In this year the Single European Act was already signed, though came into effect in 1987.

Concerning the country specific results, different break dates are found. The introduction of a break in the specification reduces notably the magnitude of the coefficient. Most of these dates are close to 1986, the abovementioned date, or to 1999, the year of the creation of EMU. All in all, CCE predicts a lower effect of the euro than CUP, though in both cases this effect is small. In some cases it is negative (Greece, Italy or Ireland, for instance) but for most members the effect is found to be positive and significant.

Acknowledgements

The authors are indebted to J. LL. Carrión-i-Silvestre and Ch. Kao for providing them with the Gauss codes to implement some of the tests used in the paper. The usual disclaimer applies. E. Gómez and C. Tamarit gratefully acknowledge the financial support from the Instituto Valenciano de Investigaciones Económicas (IVIE). The authors also acknowledge the financing from *Spanish MICINN* [projects ECO2008-05908-C02-01/ECON (M. Camarero), ECO2008-05908-C02-02/ECON (E. Gómez and C. Tamarit) and SEJ2007-62081 (E. Gómez)] M. Camarero and C. Tamarit are members of INTECO research group funded by *Generalitat Valenciana,* project PROMETEO/2009/098.

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Table 3: Banerjee and Carrion (2010) BC cointegration tests

Model	Z_j^*	r	r_1
1	-26.48	3	3
2	-33.03	3	3
3	-29.10	3	3
4	-33.27	3	3
5	**-35.72**	**3**	**3**
6	-20.42	3	2
7	-33.64	3	3
8	-37.36	3	3

Note: Model selected according to BIC. The model includes a constant, a trend and a break in both components in 1986.

Table 4
Cup estimation of the long-run parameters 1967-2008

Variables	LSDV	Bai FM	CupFM	CupBC	CCE
$\boldsymbol{GDP_{ijt}}$	**1.01**	**0.66**	**0.64**	**0.63**	**0.93**
	(80.89)	(15.94)	(17.93)	(23.28)	(10.98)
$\boldsymbol{GDPcap_{ijt}}$	**-0.22**	**-0.09**	**-0.14**	**-0.16**	**0.35**
	(-9.66)	(-2.09)	(-3.80)	(-5.57)	(2.02)
RTA	**0.39**	**0.12**	**0.02**	**-0.005**	**0.39**
	(20.57)	(10.31)	(3.28)	(-1.14)	(5.52)
EMU	**-0.18**	**-0.03**	**0.07**	**-0.03**	-0.21
	(-9.54)	(-2.16)	(6.35)	(4.64)	(-2.55)

Notes: Bold letters indicate significance at a 5% level. The specification 5 is estimated with 3 common factors and 4 lags. The t-statistic is reported in parenthesis. The year of the break is 1986.

Table 5
Country comparison of EMU effect

Variables	LSDV	Bai FM	CupFM	CupBC	CCE
Austria	0.10	**0.15**	**0.07**	**0.05**	**0.008**
1995	(1.53)	(6.01)	(3.76)	(2.50)	(0.12)
Belgium	**-0.13**	0.01	0.02	**0.11**	**0.009**
1986	(-2.30)	(0.42)	(0.91)	(5.94)	(0.35)
Finland	-0.02	**0.09**	**0.37**	**0.24**	**-0.01**
1996	(-0.41)	(3.49)	(10.01)	(8.50)	(-0.19)
France	**-0.21**	**-0.11**	**0.14**	**0.07**	**-0.033**
1991	(-4.48)	(-3.75)	(7.46)	(3.76)	(-0.54)
Germany	**-0.14**	-0.02	**0.08**	**0.16**	**0.01**
1987	(-2.84)	(-1.31)	(5.36)	(9.80)	(0.34)
Greece	-0.13	**-0.15**	**-0.40**	**-0.24**	**0.009**
1977	(-2.37)	(-5.30)	(-12.98)	(-13.59)	(0.15)
Ireland	-0.03	-0.01	**0.15**	0.04	**-0.006**
1993	(-0.52)	(-0.40)	(5.99)	(1.66)	(-0.11)
Italy	**-0.13**	**-0.06**	**-0.20**	**-0.17**	**0.005**
1999	(-3.43)	(-2.84)	(-10.89)	(-10.47)	(0.25)
Netherlands	**-0.26**	**-0.14**	0.02	**0.09**	**0.013**
1987	(-4.39)	(-4.92)	(1.17)	(5.59)	(0.27)
Portugal	-0.05	-0.01	**0.20**	**0.19**	0.004
1994	(-1.14)	(-0.37)	(9.61)	(10.44)	(0.05)
Spain	**-0.09**	**-0.06**	**0.08**	**0.07**	**0.022**
1982	(-2.20)	(-2.49)	(2.91)	(3.28)	(0.74)

Note: The specification 5 is estimated with 3 common factors and 4 lags in all cases. The t-statistic is reported in parenthesis. The year of the break is indicated below the name of each country. Bold letters indicate significance at a 5% level.